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THE DRAFT LOTTERY AND VOLUNTARY ENLISTMENT IN THE VIETNAM ERA

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ABSTRACT

A combination of voluntary enlistment, armed forces eligibility criteria, and the failure of draftees to avoid conscription jointly determined the racial composition of the Vietnam-era armed forces. Administrative data show that men with draft lottery numbers that put them at high risk of conscription are over-represented among men who voluntarily enlisted in the military, but that the effect of the lottery on enlistment is stronger for whites than for nonwhites. Minimum Chi-Square estimates of enlistment models for the 1971 draft lottery suggest that nonwhites were more likely than whites to prefer enlistment to a civilian career. This finding appears to explain racial differences in the effect of the lottery on enlistment. Contrary to the findings of a recent congressional study, the Vietnam-era estimates presented here suggest that conscription of a relatively small number of whites and nonwhites in a manner proportional to their prevalence in the population might substantially reduce nonwhite representation in the armed forces.

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"It's nonsense to cry that you can't plan your life because of the draft. You can volunteer anytime you want to."

Lieutenant General Lewis B. Hershey,
Director of Selective Service 1941-1970

1. Introduction

The draft ended in 1973, but the standby conscription plan now in effect presumes that a large-scale conflict will require resumption of the draft (Selective Service System 1986). An important issue in discussions of the draft is whether black or low income citizens have borne an excessive share of the burden of compulsory military service. For example, one common view of the Vietnam-era draft is that "the 'fittest' -- those with background, wit, or money -- managed to escape" (Baskir and Strauss 1978, p. 6). As evidence that the burden of Vietnam-era military service fell disproportionately on minorities as well as on the poor, it is often asserted that blacks were over-represented in the military (Nalty and MacGregor 1981). Yet, until the mid 1960's, the fraction of new entrants to the military who were black was generally somewhat lower than the fraction of 18 to 21 year old men who were black. Black representation among new entrants was actually highest after the draft had ended, reaching 21% in the first full year of an All-Volunteer force (1974), and declining after that. Only in 1972 did the fraction of blacks in the military reach 11%, the approximate proportion of blacks in the population (Moskos 1977).

Voluntary enlistment, armed forces eligibility criteria, and the failure of draftees to avoid conscription jointly determined the racial composition of the Vietnam-era armed forces. This combination of multiple entry routes invalidates use of the armed forces' racial composition as prima facie evidence for or against the egalitarian character of conscription. Over-representation of blacks among new entrants during the final years of the draft may be a consequence of inability to avoid conscription. But the increasing number of black entrants also reflects the increased ability of blacks to meet armed forces eligibility criteria and, as argued below, blacks' higher valuation of military service as an alternative to civilian life.

This paper develops a modeling strategy that can be used to sort out some of the factors governing entry into the Vietnam-era armed forces. Inference is based on the Vietnam-era draft lotteries, which randomly allocated the risk of conscription. In the estimation work described below, data from the 1971 draft lottery are fit to behavioral models of Vietnam-era enlistments. The estimates identify racial differences in the value of military service while controlling for differences in the probability of conscription and the probability of meeting armed forces eligibility criteria.

The plan of the paper is as follows. Section 2 provides some background on the process of selection for military service in the early 1970's. Section 3 describes the administrative records containing data on enlistment and induction. The data show that men with lottery numbers that put them at high risk of being drafted were more likely to enlist voluntarily in the military, but that the effect of the lottery on enlistment was stronger for whites than for nonwhites. Section 4 presents results from the estimation of several models of enlistment behavior during the lottery. The results in Section 4 suggest that nonwhite men were more likely than white men to prefer enlistment to a civilian career. This finding explains racial differences in the effect of the lottery on enlistment. Section 5 shows that an expected utility model fit to data on men born in 1951 does a good job of predicting the enlistment behavior of men born in 1952. A simpler linear model does not forecast as well. Section 6 provides a summary and compares the findings presented here to those of a recent government study of social representation in the armed forces.

2. Operation of the Draft Lottery

Priority for induction into the United States armed forces during the early 1970's was determined randomly in a series of draft lotteries. Each lottery consisted of the assignment of Random Sequence Numbers (RSN's, or "draft lottery numbers") to all dates of birth in the cohort to be called for service. As the year progressed, men with the lowest RSN's were called for induction first. The Defense Department determined an official RSN ceiling once manpower needs were met; only men with lottery numbers below the ceiling could have been drafted.

Separate birth cohorts were at risk in each lottery. The 1970 lottery involved men born in the years 1944-1950, the 1971 lottery involved men born in 1951, the 1972 lottery involved men born in 1952 and so on, through 1975. No one was inducted after 1972, however, and in July of 1973 congressional authority to induct soldiers expired and was never renewed.

A few months before the year in which a cohort was slated for induction, lottery numbers were assigned to exact dates of birth in a televised drawing. Besides the lottery number, further ranking based on the draft registrants' last name was also possible. In practice, however, the Selective Service relied on the screening process employed by local draft boards and the elimination of men with deferments to select from men with low lottery numbers (Tarr 1981). According to a policy known as "channeling," the Selective Service granted deferments for a variety of educational, occupational, or family reasons. Deferred men retained the liability implied by the RSN attached to their date of birth upon expiration of the deferment.

Men were selected for induction from the draft-eligible, non-deferred "high priority pool" according to several criteria, the most important of which were the pre-induction physical and the Armed Forces Qualification Test (AFQT). In 1970, for example, half the registrants failed pre-induction examinations, and the physical inspection at the time of induction eliminated a further 20% of the remaining eligible registrants (Selective Service System 1971a, pp. 5-6).

An important contribution of the draft to military manpower procurement was the production of "draft-motivated volunteers" who enlisted to avoid induction under less favorable circumstances (House Armed Services Committee 1970, Tarr 1981). Because of draft-motivated enlistments, there are large differences in probabilities of enlistment by Random Sequence Number even in cohorts from which there were few inductions (see, e.g., Figures 1 and 2 below). Binkin and Johnston (1973, Table 9) estimate that draft-motivated volunteers constituted approximately 42 percent of 1970 enlistments, 24 percent of 1971 enlistments, and 15 percent of 1972 enlistments.

Table 1 presents a chronology of events connected with the draft lottery. One important aspect of the lottery documented in this table is the uncertainty faced by young men after the assignment of random sequence numbers. Although official induction ceilings were often reached midway through the year, the highest RSN that would be called was generally not known until later in the year. For example, the 1971 ceiling for induction (125) was first called in May. But 125 was not officially declared to be the induction ceiling until October, and men with numbers as high as 170 were called for pre-induction examinations. Similarly, although no one was inducted after 1972, men born in 1953 with lottery numbers below 100 were eligible (but not necessarily called) for administrative processing in 1973.

3. Data and Descriptive Analysis

Data on enlistment and inductions come from Defense Manpower Data Center (DMDC) administrative records on entry to the military by men without prior military service. Each record in the data set provided by the DMDC contains information on the number of new entrants for a single date of birth, race and year. Draft lottery numbers were matched from tables in Selective Service System Semiannual Reports (1969-73).

The DMDC data distinguish between two types of entry to the armed forces. These are

Enlistments: Men who volunteered for military service under the terms of service available to an "enlisted volunteer." Terms of service as an enlistee usually included a 3 or 4 year tour-of-duty, choice of branch of service, the possibility of entry into specialty training or enlisted officer candidate programs, and choice of time of entry into service (House Armed Services Committee 1970, p. 12624).

Inductions: Men who either were "drafted," or volunteered for service under the same terms as draftees. Terms of service as a draftee usually included a 2 year tour-of-duty and non-specialized service in the Army.

Table 2 reports the number of new entrants by race. For each cohort born from 1951 to 1953, the table shows the number of total entrants and inductions in the years preceding and including the year in which members of the cohort were to be drafted. Data for cohorts born earlier than 1951 are not reported because the DMDC computerized records do not begin until July 1970. Also shown is the percentage of new entrants who were nonwhite. For example, 194,131 men born in 1951 entered the service between July 1970 and the end of 1971; roughly 16% of these were nonwhite. A total of 55,842 men born in 1951 were inducted between 1970 and 1971; roughly 17% were nonwhite.

Entries in each row of the table show that the fraction of nonwhite entrants rises over the sample period to a peak in 1973, the first year of the All-Volunteer Force. The fraction of nonwhite entrants and inductions by men born in 1951 and 1952 during their years of prime exposure to the draft (diagonal entries in the table) range from 14 to 17 percent. This may be compared to the 14 percent of 19 year old male citizens in 1970 who were nonwhite (U.S. Bureau of the Census 1971, Table 194).

Table 3 contains a tabulation of new entrants classified by type of entry and by draft-eligibility status. Men categorized as "draft-eligible" had lottery numbers below the highest number called in the year in which members of their cohort were to be inducted. The fraction of enlistments made by draft-eligible men documents the importance of draft-motivated enlistment. For example, in 1951 the highest lottery number called was 125; therefore 34% (125/365) of men born in 1951 were draft-eligible. Yet, 54% of the 1971 enlistees born in 1951 were draft-eligible.

The distribution of lottery numbers among enlistees also demonstrates the racial differences in draft-motivated enlistment. Figure 1a shows this distribution for white and non-white men born in 1951. Plotted in the figure are the number of volunteers for each group of 5 consecutive lottery numbers, divided by the total number of volunteers (The reason for the wavy patterns in the figure is unknown). Because lottery numbers are approximately uniformly distributed, ratios of the plotted probabilities give the relative risk of enlistment or induction by lottery number. The figure shows that whites with low lottery numbers were almost 3 times more likely to enlist than whites with high numbers. In contrast, nonwhites with low numbers were less than twice as likely to enlist as nonwhites with high numbers.

Figure 1b shows the distribution of lottery numbers among inducted men (draftees) born in 1951. The probability of induction falls to near zero for men with lottery numbers above the induction ceiling of 125. Inductions are not exactly zero for men with numbers above 125 because some men with high numbers chose to enter the military under the same terms of service as draftees. Figures 2a and 2b show the distribution of lottery numbers among enlisted and inducted men born in 1952. The enlistment response of nonwhites born in 1952 is also dampened relative to that of whites. Data for men born in 1953, not shown here, exhibit a similar pattern.

The next section describes a number of simple models of enlistment behavior in a draft lottery. These models attribute the dampened enlistment response of nonwhites to a higher propensity to consider military service an attractive alternative to the civilian labor market.

4. Enlistment During the Draft Lottery

4.1 An Expected Utility Model

A key feature of the draft lottery is the uncertainty about the prospects of induction faced by draft registrants. One way to model this uncertainty is to assume registrants (indexed by n) maximize expected utility by choosing either:

(1) to enlist and receive utility ω_n^e;
to be inducted and receive utility ω_n^i; or
to do neither and receive expected utility $q_n\omega_n^i + (1-q_n)\ \omega_n^c$,

where ω_n^c is the utility of a civilian career, and q_n is the nth man's subjective probability of being drafted. Note that men could voluntarily enlist for service under the same terms as draftees, although only a few did so.

The nth man is assumed to want to enlist in the armed forces if

$$\omega_n^e > \omega_n^i \text{ and } \omega_n^e > [q_n\omega_n^i + (1-q_n)\omega_n^c]. \qquad (2)$$

This formulation is similar to that used by Warner and Goldberg (1984) to model the effect of sea-duty on Navy reenlistment. A parsimonious statistical model is developed by letting ω_n^e, ω_n^i and ω_n^c be jointly normally distributed random variables with means M^e, M^i and M^c in the following components of variance scheme:

(3)
$$\begin{aligned} \omega_n^e &= M^e + f_n + g_n^e \\ \omega_n^i &= M^i + f_n + g_n^i \\ \omega_n^c &= M^c + f_n + g_n^c, \end{aligned}$$

where f_n is a person-specific random component with mean zero that is not a function of choice, and g_n^e, g_n^i, and g_n^c are mutually orthogonal random components with mean zero and common variance σ^2. The components of variance scheme is introduced to parameterize and restrict the correlation between the ω's for a given individual.

It is useful to transform the utility levels to the difference in utility between voluntary enlistment and civilian alternatives:

$$a_n^e = \omega_n^e - \omega_n^c = M^e - M^c + (g_n^e - g_n^c),$$

and to the difference in utility between induction and civilian alternatives:

$$a_n^i = \omega_n^i - \omega_n^c = M^i - M^c + (g_n^i - g_n^c).$$

Writing the enlistment rule in terms of utility differences, we have

(4)
$$a_n^e > a_n^i \text{ and } a_n^e > a_n^i q_n,$$

where

$$\begin{pmatrix} a_n^v \\ a_n^i \end{pmatrix} \sim N\left[\begin{pmatrix} \mu^e \\ \mu^i \end{pmatrix}, \begin{pmatrix} 2\sigma^2 & \sigma^2 \\ \sigma^2 & 2\sigma^2 \end{pmatrix}\right].$$

The parameters $\mu^e = M^e - M^c$ and $\mu^i = M^i - M^c$ are the means of a_n^e and a_n^i, and the correlation between a_n^e and a_n^i is 1/2. The estimable parameters of this distribution are $\tilde{\mu}^i = \mu^i/(\sqrt{2}\sigma)$ and $\tilde{\mu}^e = \mu^e/(\sqrt{2}\sigma)$, the standardized mean subjective costs of military service as an inducted or enlisted entrant.

Enlistment decision rule (4) is relevant for the population at risk of induction. In the Vietnam era, the population at risk was the pool of men with draft classification I-A and I-A-O (available for service) at the time draft calls began. Men classified as available for service met armed forces mental and physical standards, and were not deferred for educational or other reasons. A key identifying assumption invoked here is that the likelihood of deferment is independent of lottery numbers and of stochastic components in the enlistment decision rule. As noted in Section 2, the majority of deferments were due to failure to satisfy armed forces entry criteria. For example, of the roughly 13 million men deferred as of June 30, 1971, half did not meet the entry standards. Almost 4 million were deferred by virtue of fatherhood or other dependency, and only 1.3 million had student deferments (Selective Service System 1971b, p. 57). Graduate student deferments were eliminated in 1967 and undergraduate deferments were eliminated by late 1971, so that most of the physically and mentally acceptable men born in 1951 were eventually at risk of being drafted.

The variance-components specification provides a statistical rationale for the independence of stochastic components of the enlistment decision rule from armed forces eligibility criteria. Suppose that armed forces eligibility criteria can be described by an index, a_n^z, where

$$a_n^z = \mu^z + cf_n + g_n^z. \tag{5}$$

A draft registrant meets the eligibility criteria if a_n^z exceeds some threshold

c^*. In this expression, μ^z is the mean of a_n^z, f_n is the same as in (3), c is a constant and g_n^z is a random error component independent of g_n^c, g_n^e and g_n^i. Because f_n is the only random component common to the enlistment and eligibility decision rules, eligibility is independent of desired enlistment.

Let A_n denote the event "acceptable for service" and let W_n denote the event "wants to enlist." Then the probability of <u>successful</u> enlistment (E_n) by a man with characteristics X_n and lottery number R_n can be written

$$\begin{aligned} P(E_n | R_n, X_n) &= P(W_n, A_n | R_n, X_n) \\ &= P(W_n | R_n, X_n) P(A_n | X_n). \end{aligned}$$

To turn this into a probability statement based on (4) and (5), an additional transformation can be made to

$$b_n^1 = a_n^e - a_n^i \quad \text{and} \quad b_n^2(q_n) = a_n^e - a_n^i q_n,$$

with covariance matrix

$$\begin{bmatrix} 2\sigma^2 & (1 + q_n)\sigma^2 \\ (1 + q_n)\sigma^2 & 2(1 + q_n^2 - q_n)\sigma^2 \end{bmatrix}.$$

The probability of successful enlistment is then

$$\begin{aligned} (6) \quad & P(E_n | R_n, X_n) \\ &= \Phi[-\bar{\mu}^i + \bar{\mu}^e, -(q_n/\sqrt{(1+q_n^2-q_n)})\bar{\mu}_i + (1/\sqrt{(1+q_n^2-q_n)})\bar{\mu}^e; \rho(q_n)]P[a_n^z > c^* | X_n], \\ &= F_1[X_n, q_n]P[a_n^z > c^* | X_n], \end{aligned}$$

where $\Phi[u, v; \rho]$ is the bivariate normal distribution function with correlation coefficient ρ, and $\rho(q_n) = (1/2)[(1+q_n)/(\sqrt{1+q_n^2-q_n})]$. Note that the μ's are to be allowed to depend on X_n.

4.2 Empirical Strategy

The enlistment model parameterizes $P(E_n | R_n, X_n)$, but armed forces administrative records can only be used to estimate $P(R_n | E_n, X_n)$. However, using the definition of conditional probability and the random assignment of R_n,

$$P(E_n | R_n, X_n) = P(R_n | X_n, E_n) * [P(E_n | X_n)/P(R_n)],$$

so that parameters of $F_1(X_n, q_n)$ can be estimated by fitting

$$P(R_n|E_n, X_n) = \{F_1[X_n, q_n]P(R_n)\} * \{P(A_n|X_n)/P(E_n|X_n)\}.$$

Using (6), the marginal probability of enlistment given X_n, $P(E_n | X_n)$, can also be expressed in terms of F_1:

$$P(E_n | X_n) = P(A_n | X_n) * \sum_{j=1}^{J} F_1[X_n, q_{jn}]P(R_j),$$

where q_{nj} is the probability of induction for a man with the same characteristics (X_n) as n, but with lottery number j = 1, . . ., J. Therefore,

(7) $$P(R_n|E_n, X_n) = \{F_1[X_n, q_n]P(R_n)\}/\sum_{j=1}^{J} F_1[X_n, q_{nj}]P(R_j),$$

In the empirical work, X_n is a dummy variable that indicates race.

The remaining prerequisite for estimation is a specification for q_n, the subjective probability of induction. The simplest specification for q_n uses the empirical relative frequency of inductions by race and lottery number. The relative frequency of inductions, denoted EMP_n, is the number of inductees with characteristics X_n and lottery number R_n, divided by an estimate of the population at risk. In this case, the population at risk is the number of men classified as available for service when draft calls begin. As above, let A_n denote available status, and let NA be the number of available men in a given birth cohort. Then the population at risk with characteristics X_n and lottery number R_n is

$$P(R_n, X_n | A_n) * NA = P(R_n) * P(X_n | A_n) * NA,$$

because lottery numbers were randomly assigned. There were 682,689 men born in 1951 who were available for service on December 31, 1970 (Selective Service System 1971a, Table I).

Selective Service reports do not record the racial distribution of class I-A and I-A-O men, so that estimates of $P(X_n | A_n)$ are not available. However, information on the racial composition of inductions is available from DMDC administrative records. Therefore, I have assumed the fraction of acceptable men who were nonwhite is the same as the fraction of inducted men who were nonwhite. This is justified if the probability of being inducted, conditional on being acceptable for service, is not a function of race (Let I_n be the event "induction" and note that if $P[I_n | X_n, A_n] = P[I_n | A_n]$, then $P[X_n | I_n] = P[X_n | A_n]$). Although there are large racial differences in the probability of meeting armed forces entrance criteria (Cooper 1977), men who met these criteria were probably inducted without regard to race. In any case, except for the scaling of estimates of $\bar{\mu}^i$, the estimates were robust in experimentation with alternate denominators for q_n.

The use of EMP_n allows for variation in q_n with both race and lottery numbers, and is justified by the assumption that draft registrants beliefs' about whether they would be drafted were correct. But use of an empirical q_n does not allow for the fact that men generally had to decide to enlist before they knew whether their lottery number would be below the induction ceiling. I therefore compute additional estimates using models that characterize beliefs about the induction ceiling with a parametric distribution.

Suppose that beliefs about the induction ceiling may be described by a distribution, G, with parameters α and β. In one case below, I take G to be logistic with mean α/β and standard deviation $1/\beta$, so that the probability n believes his lottery number will be below the ceiling is

$$1 - G(R_n; \alpha, \beta) = \exp(\alpha - \beta R_n)/[1 + \exp(\alpha - \beta R_n)],$$

because the logistic distribution is symmetric. In fact symmetry may be an unrealistic assumption, if only because the induction ceiling must be positive. In an alternate specification, I take G to be a 2-parameter Weibull distribution, so that the probability n believes his number will be below the ceiling is (Johnson and Kotz 1972):

$$1 - G(R_n; \alpha, \beta) = \exp\{ -[(R_n/\alpha)^{\beta}] \}.$$

Finally, men who believe their number will be above the ceiling are assumed to believe they will not be drafted; men who believe their number will be below the ceiling are assumed to believe they will be drafted in the same proportions as were actually drafted among all draft-eligible men. Thus,

$$(8) \qquad q_n = q_n^* = QE_n * \{1 - G(R_n; \alpha, \beta)\},$$

where QE_n is the relative frequency of induction among all draft-eligible men of the same race.

The last specification I consider for q_n allows the population to contain a fraction, γ, for whom EMP_n characterizes their beliefs about being drafted, and a fraction, $(1-\gamma)$, for whom (8) characterizes their beliefs about being drafted. In this case,

$$q_n = \gamma EMP_n + (1 - \gamma)q_n^*. \tag{9}$$

The parameters α, β and γ are estimated jointly with other parameters in the model. I refer to specification (9) as a "mixed logit" or "mixed Weibull" parameterization, depending on the functional form for G. The motivation for (9) is primarily empirical; a mixture of EMP_n and a parametric specification gives a substantially better fit than either EMP_n or a parametric specification alone.

I fit the enlistment model to data on men born in 1951 who entered the military from July 1970 through the end of 1971. Estimates are for men born in 1951 because this is the only cohort from which large numbers of men were drafted, and for whom complete data on new entrants are available. The period 1970 to 1971 is used because DMDC data are unavailable before July 1, 1970 and because this is a window that begins with the lottery draw for the 1951 cohort (July 1, 1970) and includes the cohort's year of exposure to the draft.

I assume the probability of induction to be constant in groups of five consecutive lottery numbers, so that estimates may be tabulated from data grouped into 140 cells defined by race (2 values), and 5-number intervals up to lottery number 345 plus a cell for numbers 346-365 (70 values). These data correspond to the data plotted in Figure 1. The estimates minimize the following Modified Minimum Chi-Square minimand:

$$H(\theta) = \sum_{s=1}^{140} \{[\hat{p}_s - h_s(\theta)]^2/\hat{p}_s\},$$

where s indexes cells, $\hat{p}_s$ denotes the empirical $P(R_s | E_s, X_s)$, $h_s(\theta)$ denotes the right hand side of (7), and θ is the k by 1 vector of parameters. The minimization was carried out using the quadratic hill-climbing algorithm GRADX (Goldfeld and Quandt 1985). Reported standard errors are derived from the empirical second derivatives of the objective function.

When the model for the vector $\hat{p}_s$ is true, the sample size times the minimand has a chi-square distribution (Kendall and Stuart 1973). In this case, however, estimates are tabulated from a complete enumeration of the population of recruits, so that a conventional frequentist interpretation of standard errors and test statistics is not appropriate. But the test statistic provides a useful criterion for model comparison, and the standard errors indicate the degree to which each parameter contributes to the overall fit of the model.

Results of Estimation

Table 4 reports estimates of the standardized mean differences in utility between military and civilian careers ($\bar{\mu}$'s), and of the parameters in q_n (α, β and γ). Rows 1 and 2 of the table show estimates of the mean difference between the utility of induction and a civilian career, $\bar{\mu}^i$. These estimates are sensitive to the parameterization of q_n and, taken literally, imply that almost no one would choose to be drafted. An explanation for this sensitivity, however, may be the estimate of the population at risk in the denominator of q_n essentially determines the scale of $\bar{\mu}^i$. Since the true population at risk cannot be determined with certainty -- the number of non-deferred men rose and fell throughout the year -- the scale of $\bar{\mu}^i$ is probably best viewed as unidentified.

Rows 3 and 4 report estimates of the difference between the value of enlistment and a civilian career, $\bar{\mu}^e$. These estimates are not very sensitive to the parameterization of q. To interpret the parameter estimates, rows 8 and 9 of the table shows estimates of the probability of preferring enlistment to a civilian career, $P(\omega_n^e > \omega_n^c | X_n)$, computed by substituting estimates of $\bar{\mu}^e$ in the normal distribution function. Row 10 reports odds ratios that compare estimates of the probability of preferring enlistment by race.

The estimates of $\tilde{\mu}^e$ suggest that 20 to 30 percent of whites found enlistment attractive, whereas 40 to 60 percent of nonwhites found enlistment attractive. Are these numbers reasonable? Cooper (1977, Table 10-9) has estimated the number of men in the population who would meet the armed forces standards for mental ability, and then used these figures to estimate "military participation rates" by race. Cooper's figures suggest that 24% of acceptable 19 year old white men enlisted in 1971-72, while 53% of acceptable 19 year old nonwhites enlisted in the same period. Thus, results from the expected utility model are roughly consistent with Cooper's estimated participation rates. Except for an even larger odds ratio in column (1), the ratios in row 10 consistently suggest that the odds of a nonwhite man preferring enlistment were 2 to 3 times larger than the odds of a white man preferring enlistment.

Estimates of α, β and γ are reported in rows 5, 6 and 7 of Table 4. The mean of the logistic prior distribution on the induction ceiling is $-5\alpha/\beta$ (the table reports the negative of β). For the logit specifications, estimates of α and β reported in column 2 imply an expected induction ceiling of RSN 120, and estimates in column 4 imply an expected induction ceiling of RSN 171. For the Weibull specifications, estimates of β in column 3 imply a skewed distribution with an extended right tail, and estimates in column 5 imply a skewed distribution with an extended left tail. An approximation to the mean of a two-parameter Weibull distribution is given by Johnson and Kotz (1972, p. 253). Using this approximation, the Weibull ceilings are estimated to be around RSN 111 in column (3) and RSN 163 in column (5). Note that the estimated ceilings for the mixed logit and Weibull specifications are between the 1970 ceiling of 195 and the actual 1971 ceiling of 125 that applied to the 1951 cohort.

4.3 A Linear Model

An unattractive feature of the expected utility model is the need for an unverifiable distributional assumption to characterize random utilities. A model with fewer distributional assumptions can be developed by assuming that draft registrants believe they know whether they will be drafted. Let d_n be a dummy variable that equals one if registrant n expects to be drafted, and let $q_n = P(d_n = 1| X_n, R_n)$. In this formulation, the probability that n would like to enlist given X_n and R_n is simply

$$F_2(X_n, q_n) = P(\omega_n^e > \omega_n^i | X_n)q_n + P(\omega_n^e > \omega_n^i, \omega_n^e > \omega_n^c | X_n)(1 - q_n).$$

In contrast to the expected utility model, the assumption that men believe they know whether they will be drafted leads to a probability statement that is linear in q_n. Note, however, that the interpretation of q_n has changed. In the expected utility model, men are uncertain about whether they will be drafted; q_n is a subjective degree of belief held by the nth man. Here, perhaps less realistically, men decide whether they will be drafted; q_n is the frequentist probability that a man chosen at random with characteristics X_n and lottery number R_n is one who believes he will be drafted.

As in the expected utility model, armed forces eligibility criteria are assumed to be independent of $\omega_n^e - \omega_n^i$ and $\omega_n^e - \omega_n^c$, and independent of lottery numbers. Therefore the probability of successful enlistment given X_n and R_n is

$$P(E_n| X_n, R_n) = F_2(X_n, q_n)P(A_n| X_n).$$

The assumption that draft registrants behave as if they know whether they will be drafted leads to a linear model that can be used to identify the ratio of

$P(\omega_n^e > \omega_n^i,\ \omega_n^e > \omega_n^c \mid X_n)$ to $P(\omega_n^e > \omega_n^i \mid X_n)$. To see this, write

$$F_2(X_n,\ q_n) = \pi_0 q_n + \pi_1(1-q_n),$$

where

$$\pi_0 = P(\omega_n^e > \omega_n^i \mid X_n),$$
$$\pi_1 = P(\omega_n^e > \omega_n^i,\ \omega_n^e > \omega_n^c \mid X_n)].$$

The dependence of π_0 and π_1 on X_n is suppressed for notational clarity. Substituting $F_2(X_n,\ q_n)$ for $F_1(X_n,\ q_n)$ in (7), and defining $P(I_n \mid X_n) = \sum_{j=1}^{J} q_{nj} P(R_j)$ we have,

$$P(R_n \mid E_n,\ X_n) = \{\delta_0 q_n + \delta_1[1 - q_n]\} P(R_n),$$

where

$$\delta_0 = \pi_0/[\pi_1 + (\pi_0-\pi_1)P(I_n \mid X_n)],$$
$$\delta_1 = \pi_1/[\pi_1 + (\pi_0-\pi_1)P(I_n \mid X_n)].$$

Thus, π_1/π_0 is identified from the ratio of unrestricted estimates of δ_1 and δ_0, although π_1 and π_0 are not separately identified. Note that π_1/π_0 can be taken as an approximation to $P(\omega_n^e > \omega_n^c \mid X_n)$. This is only an approximation because $\omega_n^e - \omega_n^i$ and $\omega_n^e - \omega_n^c$ are clearly not independent in the variance components formulation (3). But if $P(\omega^e > \omega^i)$ is close to 1, as suggested by estimates from the expected utility model, the approximation should be good. Moreover, when q_n is not a function of unknown parameters (i.e., $q_n = EMP_n$), the estimation strategy requires no parametric distributional assumptions and reduces to weighted least squares in a linear model.

An interesting difference between the linear and expected utility models is in the ability of the two models to generate theoretical predictions about the effect of the draft. The effect of q_n on enlistment in the linear model is

$$P(\omega^e > \omega^i) - P(\omega^e > \omega^i, \omega^e > \omega^c).$$

Clearly, other things equal, groups valuing enlistment more (for whom ω^e is more likely to exceed ω^c) will be less likely to enlist in response to a draft. In contrast, in the expected utility model, the relationship between $\omega^e - \omega^c$ and the effect of the draft on enlistment is theoretically ambiguous.

Results of Estimation

Estimates of the reduced-form parameters δ_0 and δ_1 are tabulated by substituting F_2 for F_1 in (7) and fitting (7) to the relative frequencies plotted in Figure 1. Table 5 reports the results of this estimation. Here, the estimates are less sensitive to the choice of specification for q_n than in the expected utility model. Estimates of δ_0 are around 5.5 for whites and 3.3 for nonwhites. Estimates of δ_1 are around .55 for whites and .77 for nonwhites. The effect of an increase in the fraction who believe they will be drafted on the fraction who want to enlist is proportional to $\delta_0 - \delta_1$. In each specification, this effect is substantially larger for whites than for nonwhites.

Table 5 also shows estimates of π_1/π_0, computed as δ_1/δ_0. These estimates range from 10 to 12 percent for whites and 22 to 26 percent for nonwhites. For each race, the magnitudes of the fraction estimated to find the military attractive is roughly half that estimated in the expected utility model. But estimates from the linear model replicate the finding from the expected utility model that the odds of preferring enlistment to a civilian career are nearly 3 times larger for nonwhites than for whites.

A simple way to measure the scale of racial differences in the impact of the

the draft on enlistment is with logarithmic derivatives -- the percentage increase in the probability of enlisting in response to a one percentage point increase in the fraction who think they will be drafted. For the linear model, this is

$$[1 - (\pi_1/\pi_0)]/[q_n + (\pi_1/\pi_0)(1-q_n)].$$

Using parameters from the mixed logit specification and evaluating this expression at the probability of being drafted for draft-eligible men (.213 for whites and .202 for nonwhites), a one percentage point increase in the fraction who think they will be drafted leads to about a 3 percent increase in enlistment for whites and about a 2 percent increase in enlistment for nonwhites. Other things equal, the estimates suggest that conscription of equal proportions of whites and nonwhites in the Vietnam era would have reduced nonwhite representation in the armed forces.

5. Goodness-of-fit and Out-of-Sample Prediction

Figures 3a and 3b show the actual and fitted values for the expected utility and linear models estimated using the mixed logit specification of q_n. The figures show that both models fit the data well for both races. Note that the unrestricted γ in the mixture specification of q_n (Column (4) in Tables 4 and 5) allows for a discrete jump in the estimated probability of enlistment near the induction ceiling of RSN 125.

The last rows of Tables 4 and 5 report chi-square goodness-of-fit statistics, equal to the minimized objective function. When compared with conventional 1% critical values, each chi-square test results in a decisive rejection of the models. The critical values are roughly 170, while the smallest chi-square statistic is 864 (for the linear model estimated with mixed logit q_n.) But Figure 3 and the large number of observations (approximately 138,000) suggest that critical values for a classical test of fixed size may not provide a good standard for comparison. As originally pointed out by Berkson (1938), with enough observations, any point null hypothesis will be rejected.

An alternative to goodness-of-fit criteria for model evaluation is the quality of out-of-sample predictions. If the estimates in Tables 4 and 5 really capture behavioral parameters, then these estimates should be useful for predicting the behavior of men born in 1952, who participated in the 1972 draft lottery. But estimates of the parameters underlying q_n from the 1951 data are unlikely to be useful for predicting the probability of induction in the 1952 lottery because draft registrants in the lottery period were keenly aware that the war was winding down and that fewer and fewer men were to be called for service (Tarr 1981). The part of the model that should be invariant from year to year is the characterization of racial differences in enlistment behavior.

To compute out-of-sample predictions, the 1951 estimates of all parameters except α, β, and γ were entered as constants in models fit to data on the men born in 1952 who entered the military in 1971 and 1972. One set of estimates was computed treating the 1951 estimates of $\bar{\mu}^e$ and $\bar{\mu}^i$ as constants in the expected utility model, and a second set was computed treating the 1951 estimates of δ_0 and δ_1 as constants in the linear model. In both cases, estimates using the mixed logit specification for q_n are reported. The empirical components of q_n (EMP_n and QE_n) were constructed in a manner analogous to that used for 1951 and are reported in the notes to Table 6, which also reports the resulting estimates of α, β, and γ.

The mixed-logit expected utility model, with a chi-square of 1393, clearly does a better job of predicting the 1952 data the corresponding linear model, with a chi-square of 2056 (the expected utility model out-performed the linear model regardless of the specification for q_n). This finding is illustrated in Figures 4a and 4b, which show actual and predicted values from the expected utility and linear models. Figure 4a shows that the expected utility model does a remarkably good job of predicting the 1952 data with 1951 parameter estimates. In contrast, Figure 4b shows that the linear model under-predicts the representation of draft-eligible men among enlistees. The estimates of α and β in Table 6 give further evidence against the linear model: the estimated RSN ceiling of 57 for the expected utility model is much closer to the actual ceiling of 95 for the 1972 lottery than the estimated ceiling of 349 for the linear model. On the other hand, both models succeed in capturing the reduced impact on the draft on nonwhite enlistments.

6. Summary and Conclusions

Administrative data on enlistment and induction in the Vietnam Era show that whites were more likely than nonwhites to have enlisted voluntarily in response to the risk of being drafted. The influence of factors affecting the decision to enlist is estimated here by exploiting the random assignment of the risk of induction in the 1971 draft lottery. Racial differences in response to the lottery appear to be explained by the fact that nonwhites were more likely than whites to have considered enlistment an attractive alternative to a civilian career. Both an expected utility model and a simpler linear model suggest that nonwhites were roughly twice as likely as whites to have preferred enlistment in the military to a civilian career. This race difference is consistent with results I have presented elsewhere (Angrist 1990) which suggest that white Vietnam veterans earn 15 percent less than nonveterans as much as 10 years after their discharge from service, while nonwhite veterans suffer no earnings loss.

Estimates from the best-fitting expected utility model imply that 26 percent of whites who were available for military service found the military attractive, while 41 percent of available nonwhites found the military attractive. Estimates from the corresponding linear model imply that 10 percent of available whites preferred enlistment, while 23 percent of available nonwhites preferred the military. In an attempt to compare and evaluate the two models, estimates for men born in 1951 were used to predict the distribution of draft lottery numbers among men born in 1952. The expected utility model does a remarkably good job of predicting the distribution of lottery numbers among enlistees from the younger cohort, while the linear model under-predicts the fraction of enlistees who were draft-eligible. Estimates from the expected utility model are also closer to independently produced estimates of the fraction of the military manpower pool who enlisted.

An implication of the results from both models is that demographic groups that are most likely to view military service as unattractive are also most likely to enlist in response to a draft. This appears to explain why whites are more likely to enlist in response to a draft than nonwhites. Advocates of a return to conscription (e.g. Fallows 1989, Lacy 1982) rarely propose a return to the controversial deferment policies of the sixties and seventies as well. Estimates from the Vietnam era suggest that, without the possibility of deferment, the induction of whites and nonwhites in a manner proportional to their prevalence in the population may generate proportionately more white than nonwhite enlistments. The estimates presented here imply that the Vietnam-era draft generated a 3 percent increase in white enlistment for every 2 percent increase in nonwhite enlistment.

The findings in this paper contrast sharply with those of a recent government study (U.S. Congress 1989) on the effects of a draft on social representation in the armed forces. The study argues that a new draft would do little to improve the representation of currently under-represented groups because (page 13):

". . . if conscription was reinstated, not more than one in four of the males reaching enlistment age each year would need to serve, and unless volunteering was sharply curtailed less than 10 percent of males would be compelled to serve. Thus, conscription alone would not ensure that many of the country's future leaders would be called upon to serve . . ."

But if the results presented here for the Vietnam era generalize, then even with relatively low probabilities of conscription, draft-motivated enlistment will be a powerful force that increases the representation of groups who would not otherwise choose to enter the armed forces. Roughly one-third of Vietnam-era cohorts served in the military, and roughly one third of Vietnam-era veterans were drafted, so that the Vietnam experience is not very different from the hypothetical draft envisioned in the congressional study. Of course, one reason for draft-motivated enlistment is to reduce the risk of combat from service in the army, so that draft-motivated enlistment may be less important in a peacetime draft than it was during the Vietnam war. But resumption of the draft is most likely in the event of a large scale conflict, in which case the impact of draft-motivated enlistment on social representation in the military may equal or even exceed the impact of draft-motivated enlistment on the Vietnam-era armed forces.

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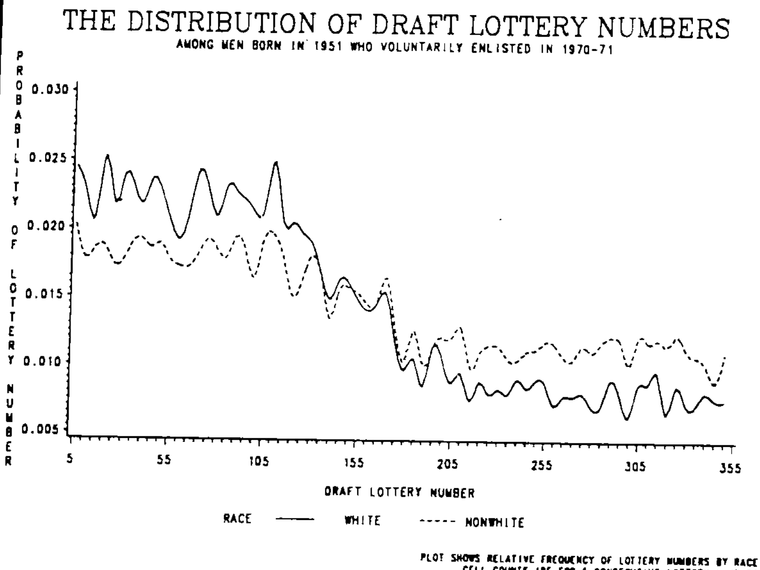

Figure 1a

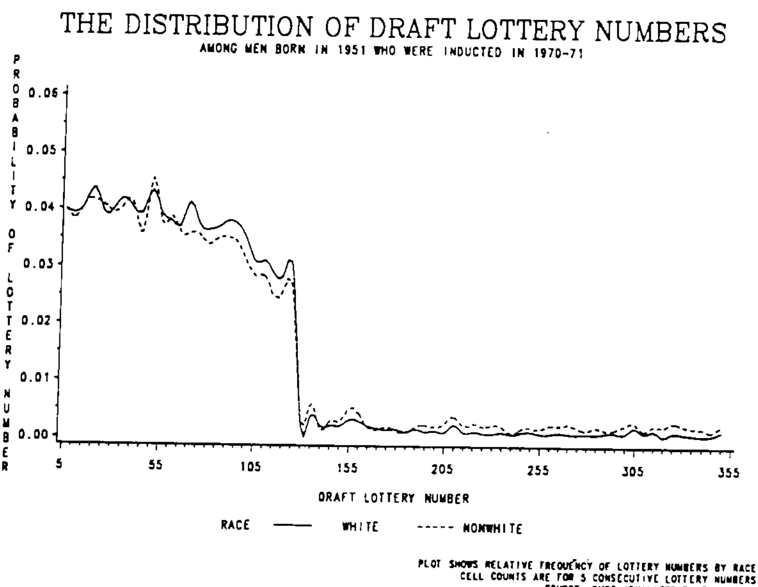

Figure 1b

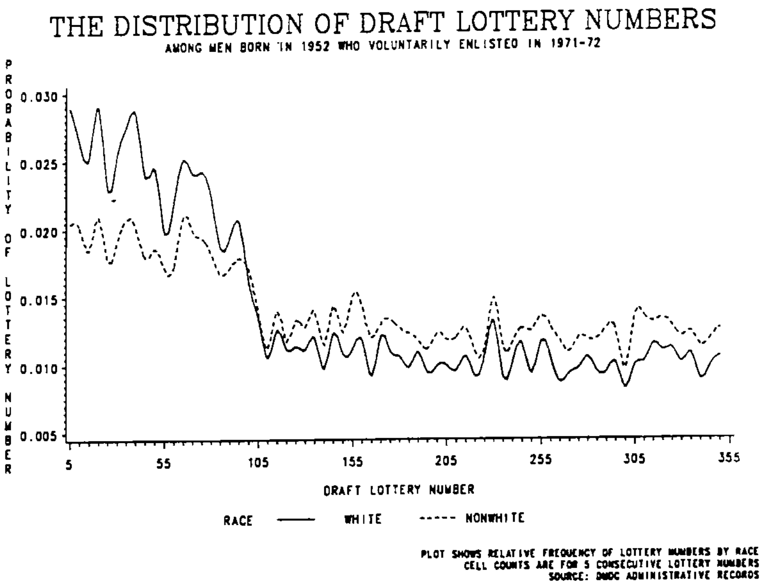

Figure 2a

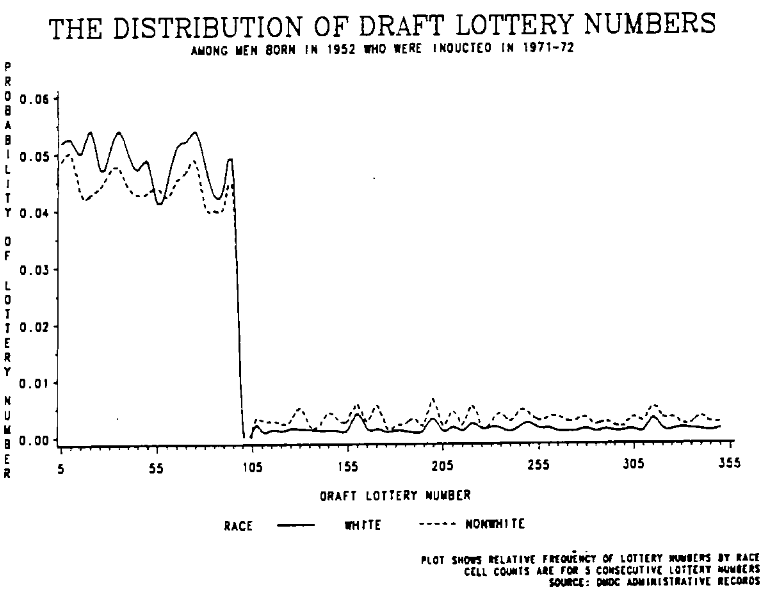

Figure 2b

ACTUAL AND FITTED VALUES FROM EXPECTED UTILITY MODEL

MEN BORN IN 1951 WHO VOLUNTARILY ENLISTED IN 1970-71

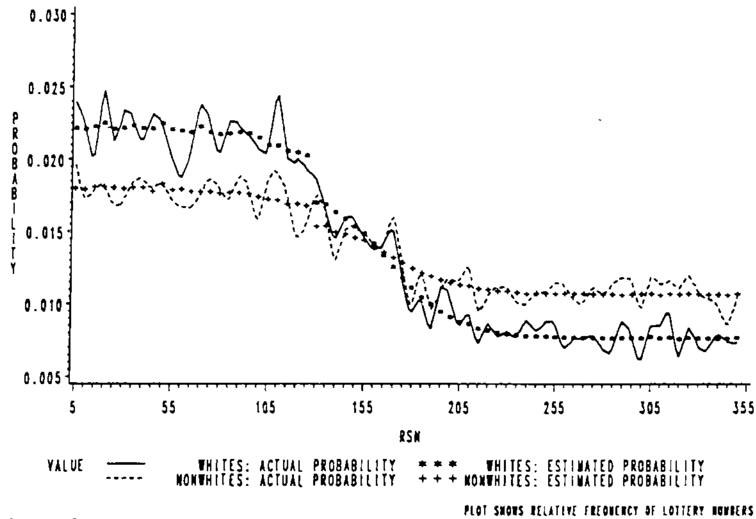

Figure 3a

ACTUAL AND FITTED VALUES FROM LINEAR ENLISTMENT MODEL

MEN BORN IN 1951 WHO VOLUNTARILY ENLISTED IN 1970-71

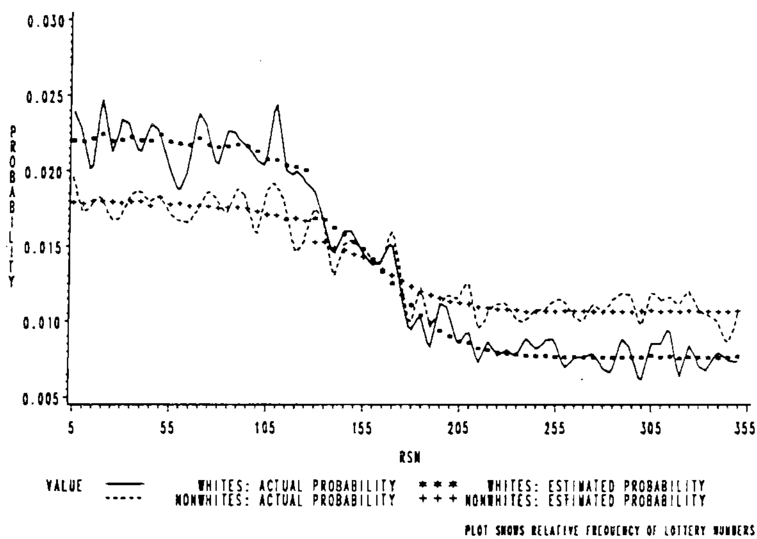

Figure 3b

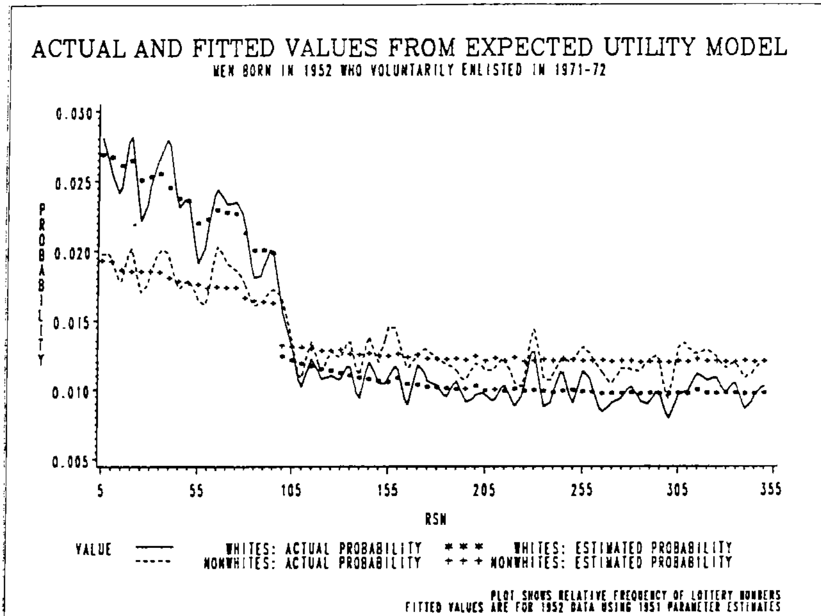

Figure 4a

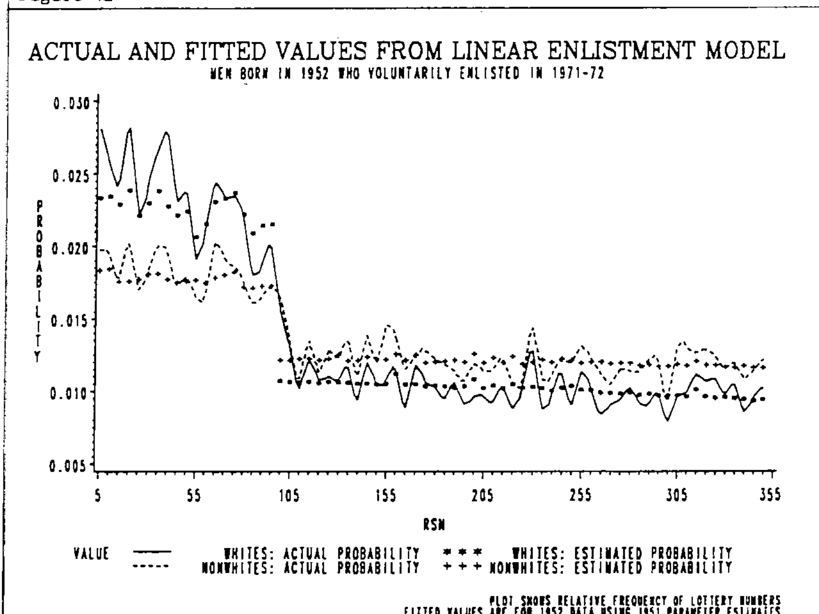

Figure 4b

Table 1

A Chronology of Random Selection

Lottery	Event	Date
1970	Random Selection for Birth Cohorts 1944-50	December 1, 1969
	RSN 1-30 Called	January 1970
	" 1-60 "	February
	" 1-90 "	March
	" 1-115 "	April
	" 1-145 "	May
	" 1-170 "	June
	" 1-190 "	July
	" 1-195 "	August - December
	CEILING: 195	
1971	Random Selection for Birth Cohort 1951	July 1, 1970
	RSN 1-100 Called	January - April 1971
	" 1-125 "	May - December
	CEILING ANNOUNCED: 125	October
1972	Random Selection for Birth Cohort 1952	August 5, 1971
	Secretary of Defense Announces No Draft Calls for January - March 1972	January 30, 1972
	RSN 1-15 Called	April - June 1972
	" 1-75 "	September
	" 1-95 "	October - December
	CEILING ANNOUNCED: 95	September

Table 1 (cont.)

1973	Random Selection for Birth Cohort 1953	February 2, 1972
	Secretary of Defense Announces No Draft Calls for January, 1973	November 28, 1972
	Secretary of Defense Announces No Further Draft Calls Foreseen	January 27, 1973
	Processing, Examination and Induction Suspended (Some Processing of Low Numbers Continued)	February 14, 1973
	Processing of Suspended Inductions Resumes	April 4, 1973
	Orders Issued to terminate all Inductions After July 1, 1973	May 15, 1973
	Induction Authority Expires (for those never deferred)	July 1, 1973
	NO FURTHER INDUCTIONS	

Note: In 1973, registrants with numbers below 100 were eligible for administrative processing. In 1974, the ceiling for administrative processing was 95. The last lottery was held in 1975, for those born in 1956.

Sources: Selective Service System (1969-73, 1986).

Table 2

New Entrants to the Military by Year of Birth and Year of Entry

Cohort	Years of entry into service					
	1970-71	Percent nonwhite	1971-72	Percent nonwhite	1972-73	Percent nonwhite
			All new entrants			
1951	194,131	15.8	170,587	17.5	44,767	24.9
1952	166,551	14.9	240,135	16.1	154,020	17.9
1953	118,954	13.3	214,659	16.2	168,490	19.7
Total	479,636	14.9	625,381	16.5	367,277	19.6
			New inductions (draftees)			
1951	55,842	16.6	55,150	16.6	2,332	19.0
1952	4,048	25.0	40,355	14.2	38,009	13.4
1953	933	25.5	1,599	28.3	846	28.5
Total	60,823	17.3	97,104	15.8	41,187	14.0

Notes: 1970 entrants from July - December only.

Source: Defense Manpower Data Center administrative records.

Table 3

New Entrants to the Military by Draft-Eligibility Status

Cohort	Year	Draft-eligible	Type of Entry			
			Enlistment	Percent	Induction	Percent
1951	1970	no	25312		1409	
		yes	24937	49.6	1562	52.6
	1971	no	40129		2998	
		yes	47902	54.4	49864	94.3
	1972	no	17803		299	
		yes	9596	35.0	1981	86.9
	1973	no	10904		4	
		yes	4131	27.5	48	92.3
1952	1970	no	37930		1107	
		yes	13285	25.9	332	23.1
	1971	no	72365		1805	
		yes	38913	35.0	804	30.8
	1972	no	41875		1678	
		yes	46621	52.7	36068	95.6
	1973	no	20269		12	
		yes	7244	26.3	251	95.4
1953	1970	no	13568		134	
		yes	5121	27.4	44	24.7
	1971	no	73458		543	
		yes	25873	26.1	212	28.1
	1972	no	74261		485	
		yes	39465	34.7	359	42.5
	1973	no	37768		1	
		yes	16148	30.0	1	50.0

Notes: Eligibility status is defined by the official RSN ceiling for the 1951 and 1952 cohorts and is set at RSN 95 for the 1953 cohort.

1970 entrants from July - December only.

Source: Defence Manpower Data Center administrative records.

Table 4

Estimates of the Expected Utility Model

Parameter	Functional Form for Probability of Induction (q_n)				
	Empirical (1)	Logit (2)	Weibull (3)	Mixed Logit (4)	Mixed Weibull (5)
1. $\bar{\mu}^i$ - whites	-69.8 (1.58)	-15.4 (4.05)	-22.9 (6.70)	-5.96 (1.51)	-7.80 (3.52)
2. $\bar{\mu}^i$ - nonwhites	-26.1 (3.69)	-11.0 (3.30)	-16.7 (4.68)	-3.35 (1.67)	-5.04 (1.63)
3. $\bar{\mu}^e$ - whites	-0.781 (0.013)	-0.416 (0.009)	-0.405 (0.008)	-0.642 (0.178)	-0.496 (0.068)
4. $\bar{\mu}^e$ - nonwhites	0.033 (0.052)	0.214 (0.029)	0.230 (0.025)	-0.234 (0.477)	0.051 (0.150)
5. α		3.93 (0.75)	23.3 (2.57)	9.56 (0.98)	35.1 (0.91)
6. β		-0.164 (0.013)	2.10 (0.29)	-0.280 (0.024)	5.38 (0.82)
7. γ				0.287 (0.057)	0.289 (0.045)
8. $P(\omega^e > \omega^c)$ whites	0.218 (0.004)	0.339 (0.003)	0.343 (0.003)	0.260 (0.058)	0.310 (0.024)
9. $P(\omega^e > \omega^c)$ nonwhites	0.513 (0.021)	0.585 (0.011)	0.591 (0.010)	0.408 (0.185)	0.520 (0.060)
10. odds ratio: nonwhites/whites	3.78	2.75	2.77	1.96	2.41
11. χ^2(dof)	2424 (134)	929 (132)	925 (132)	873 (131)	889 (131)

Notes:

Standard errors in parentheses. Sample size: 21,387 nonwhites; 116,893 whites. α, β, and γ are from

$$q_n = \gamma EMP_n + (1 - \gamma) \{[1 - G(R_n|\alpha, \beta)] * QE_n\}$$

Estimated population at risk for EMP_n: 8,134 white and 1,619 nonwhite for each cell. QE_n is the average of EMP_n below the induction ceiling of 195, equal to .213 for whites, and .202 for nonwhites. There were 682,689 men born in 1951 and classified I-A and I-A-O on December 31, 1970 (Selective Service System 1971a, Table I).

Table 5

Estimates of the Linear Enlistment Model

Parameter	Functional Form for Probability of Induction (q_n)				
	Empirical (1)	Logit (2)	Weibull (3)	Mixed Logit (4)	Mixed Weibull (5)
1. δ_0-whites	5.19 (0.03)	5.58 (0.04)	5.54 (0.04)	5.44 (0.04)	5.44 (0.04)
2. δ_0-nonwhites	3.16 (0.76)	3.42 (0.08)	3.40 (0.08)	3.35 (0.08)	3.35 (0.08)
3. δ_1-whites	0.606 (0.003)	0.552 (0.004)	0.559 (0.004)	0.547 (0.004)	0.552 (0.004)
4. δ_1-nonwhites	0.802 (0.008)	0.775 (0.009)	0.778 (0.009)	0.767 (0.009)	0.771 (0.009)
5. α		6.90 (0.21)	33.2 (0.15)	10.0 (0.83)	35.4 (0.32)
6. β		-0.225 (0.007)	4.71 (0.14)	-0.296 (0.022)	6.33 (0.45)
7. γ				0.297 (0.032)	0.262 (0.031)
8. $\delta_1/\delta_0 = \pi_1/\pi_0$ whites	0.117 (0.001)	0.099 (0.001)	0.101 (0.001)	0.100 (0.001)	0.102 (0.001)
9. $\delta_1/\delta_0 = \pi_1/\pi_0$ nonwhites	0.254 (0.008)	0.227 (0.007)	0.229 (0.007)	0.229 (0.007)	0.230 (0.007)
10. odds ratio: nonwhites/whites	2.57	2.67	2.64	2.67	2.63
11. χ^2(dof)	4253 (134)	926 (132)	944 (132)	864 (131)	885(131)

Notes:

Under the assumptions given in the text, π_1/π_0 is an estimate of $P(\omega^e > \omega^c)$. See notes to Table 4.

Table 6

Enlistment Models for Men Born in 1952

Parameter	Linear Model (1)	Expected Utility Model (2)
α	3.14	1.36
	(0.22)	(0.08)
β	-0.045	-0.119
	(0.003)	(0.004)
γ	0.790	0.400
	(0.007)	(0.013)
χ^2(dof)	2056 (135)	1393 (135)

Notes:

Standard errors in parentheses.
Sample: 166,876 whites and 32,898 non-whites born in 1952 who entered the military in 1971 and 1972.
Estimates of α, β, and γ for men born in 1952 conditional on parameter estimates for men born in 1951. Mixed logit specification used for q_n in both models.
Estimates of the population at risk for EMP_n: 7,414 whites and 1,227 non-whites in each cell. QE_n is the average of EMP_n below the induction ceiling of 95, equal to .228 for whites and .206 for non-whites. There were 604,831 men born in 1952 and classified I-A and I-A-O on June 30, 1971 (Selective Service System 1971b, Table I).